Probability and Statistics for JMC

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May 17, 2021

Contents

Chapter 1

Review of Elementary Set Theory

Ω	universal set
$\emptyset$	empty set
$A \subseteq \Omega$	subset of Ω
$\overline{A}$	Complement of A
$\lvert A \rvert$	cardinality of A
$A \cup B$	union (A or B)
$A \cap B$	intersection(A and B)
$A = B$	both sets have exactly the same elements
$A \backslash B$	set difference (elements in A that are not in B)
$\{\omega\}$	a singleton with only the element ω in the set
$A \times B$	$\big\{(a,b) \vert a \in A, b \in B\big\}$

Chapter 2

Visual and Numerical Summaries

2.1 Visualization

Definition 1. The ***histogram*** allows us to visualize how a sample of data is distributed, say the observed values are $\{x_1, \ldots, x_2\}$. The first step is deciding on a set of ***bins*** that divide the range of x into a series of intervals. A histogram then shows the ***frequency*** for each bin.

<u>**Comments**</u> Often the histogram's y-axis is normalized in some way.

- Instead of showing frequency, the height of the histogram can show ***relative frequency***, the fraction of the data set contained within the bin. In this case, $1 = \sum_{\text{bins } i} y_i$, where y_i is the relative frequency at bin i.
- The histogram could also show the ***density***, the relative frequency divided by the bin width. In this case, $1 = \sum_{\text{bins } i} \rho_i \Delta x_i$, where ρ_i is the density for bin i and Δx_i is the width of bin i.

Definition 2. The ***empirical cumulative distribution function*** of a sample of real values $\{x_1, \ldots, x_n\}$ is

$$F_n(x) = \frac{1}{n} \sum_{i=1}^{n} I(x_i \leq x),$$

where $I(x_i \leq x)$ is an ***indicator function***, i.e. the value is 1 when $x_i \leq x$ and 0 when $x_i > x$.

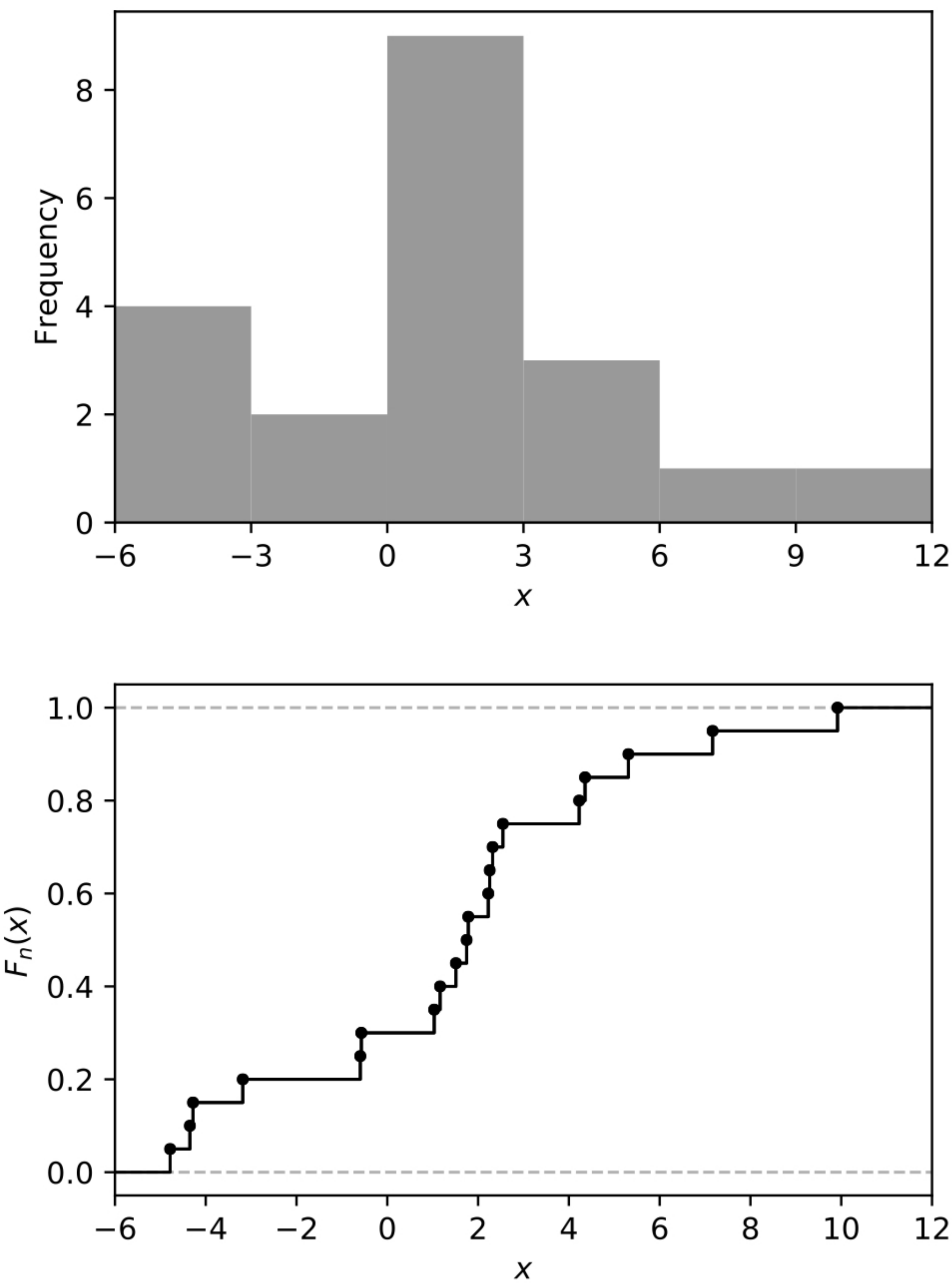

Figure 2.1: The first diagram is the histogram, and the second diagram is the empirical cdf with the same set of data

2.2 Summary Statistics

2.2.1 Measures of Location

arithmetic mean	$\overline{x} = \frac{1}{n}\sum_{i=1}^{n} x_i$
geometric mean	$x_G = \left(\prod_{i=1}^{n} x_i\right)^{\frac{1}{n}}$
harmonic mean	$\frac{1}{x_H} = \frac{1}{n}\sum_{i=1}^{n}\frac{1}{x_i}$
i^{th} order statistic	$x_{(i)}$ = the i^{th} smallest value of the sample
median	$x_{\left(\frac{n+1}{2}\right)}$
mode	x_i which occurs most frequently in the sample

Comments

- For positive data $\{x_1, \ldots, x_n\}$,

  $$\text{arithmetic mean} \geq \text{geometric mean} \geq \text{harmonic mean}.$$

- Arithmetic mean and geometric mean are related in the following way:

  $$\overline{x} = \frac{1}{n}\sum_{i=1}^{n} x_i = \frac{1}{n}\sum_{i=1}^{n}\ln y_i = \frac{1}{n}\ln\prod_{i=1}^{n} y_i = \ln\left(\prod_{i=1}^{n} y_i\right)^{\frac{1}{n}} = \ln x_G,$$

  where $x_i = \ln y_i$.

- For $x_{(i)}$, when i is not an integer, we define $\alpha \in (0,1)$ s.t. $\alpha = i - \lfloor i \rfloor$, and

  $$x_{(i)} = (1-\alpha)x_{(\lfloor i \rfloor)} + \alpha x_{(\lceil i \rceil)}.$$

2.2.2 Measures of Dispersion

mean square/sample variance	$s^2 = \frac{1}{n}\sum_{i=1}^{n}(x_i - \overline{x})^2$
root mean square/sample standard deviation	$s = \sqrt{s^2}$
range	$x_{(n)} - x_{(1)}$
first quartile	$x_{\left(\frac{1}{4}(n+1)\right)}$
third quartile	$x_{\left(\frac{3}{4}(n+1)\right)}$
interquartile range	$x_{\left(\frac{1}{4}(n+1)\right)} - x_{\left(\frac{3}{4}(n+1)\right)}$

Comments

- sample variance's different expression:

$$s^2 = \frac{1}{n}\sum_{i=1}^{n}(x_i - \overline{x})^2 = \left(\frac{1}{n}\sum_{i=1}^{n}x_i^2\right) - \left(\frac{1}{n}\sum_{i=1}^{n}x_i\right)^2 = \overline{x^2} - \overline{x}^2.$$

- Robustness, shown in table 2.1

Table 2.1: Robustness of different location and dispersion statistic

	Least Robust	More Robust	Most Robust
Location	$\frac{x_{(1)}+x_{(n)}}{2}$	$\overline{x}$	$x_{\left(\frac{n+1}{2}\right)}$
Dispersion	$x_{(n)} - x_{(1)}$	s	$x_{\left(\frac{3}{4}(n+1)\right)} - x_{\left(\frac{1}{4}(n+1)\right)}$

2.2.3 Covariance, Correlation, and Skewness

covariance	$s_{xy} = \frac{1}{n}\sum_{i=1}^{n}(x_i - \overline{x})(y_i - \overline{y})$
correlation	$r_{xy} = \frac{s_{xy}}{s_x s_y}$
skewness	$\frac{1}{n}\sum_{i=1}^{n}\left(\frac{x_i - \overline{x}}{s}\right)^3$

Comments

- covariance's different expression:

$$s_{xy} = \frac{1}{n}\sum_{i=1}^{n}(x_i - \overline{x})(y_i - \overline{y}) = \frac{1}{n}\sum_{i=1}^{n} x_i y_i + \frac{1}{n}\sum_{i=1}^{n} -x_i\overline{y} - \overline{x}y_i + \overline{xy} = \frac{\sum_{i=1}^{n} x_i y_i}{n} - \overline{x}\,\overline{y}.$$

  In the random variable's context, it is

$$\text{Cov}(X, Y) = E(XY) - E(X)E(Y).$$

- Correlation gives a **scale-invariant** measurement of relatedness between x and y, since

$$|r_{xy}| \leq 1.$$

- A sample is **positively (negatively)** or **right (left) skewed** if the upper tail of the histogram of the sample is longer (shorter) than the lower tail.

2.2.4 Box-and-whisker plot

The diagram is based on the five-point summary (use Figure 2.2 as reference):

- Median – middle line in the box.
- 3^{rd} and 1^{st} Quartiles – top and bottom of the box.
- "Whiskers" – extend out as dashed lines from the box to max/min values, which are the two short horizontal lines.
- Any outliers, i.e. extreme points beyond the whiskers, are plotted individually as dots.

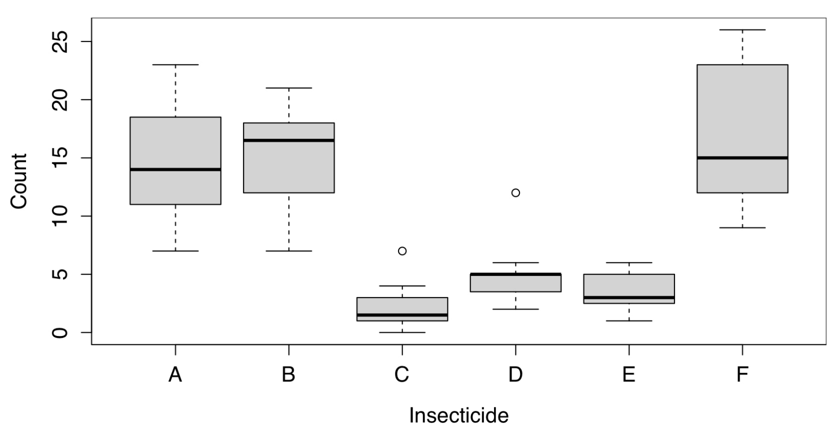

Figure 2.2: the counts of insects found in agricultural experimental units treated with six different insecticides A-F

Chapter 3

Probability

3.1 Formal Definition of Probability

3.1.1 σ-algebra

Definition 3. $\mathcal{F}$, a collection of subsets of a set S, is called a ***σ-algebra*** associated with S if:

(a) $S \in \mathcal{F}$,

(b) $\mathcal{F}$ is closed under complements w.r.t. S:

$$E \in \mathcal{F} \Longrightarrow \overline{E} \in \mathcal{F},$$

(c) $\mathcal{F}$ is closed under countable unions:

$$E_1, E_2, \ldots \in \mathcal{F} \Longrightarrow \bigcup_{i=1}^{\infty} E_i \in \mathcal{F}.$$

Comments Definition 3 implies two facts.

1. $\mathcal{F}$ must contain the empty set $\emptyset$.

   Proof. Since $S \in \mathcal{F}$, we have $\overline{S} = \emptyset \in \mathcal{F}$. □

2. $\mathcal{F}$ must be closed under countable intersections.

   Proof. Let $E_1, E_2, \ldots \in \mathcal{F}$. We can then imply the following:

$$\overline{E_1}, \overline{E_2}, \ldots \in \mathcal{F} \Rightarrow \bigcup_i \overline{E_i} \in \mathcal{F} \Rightarrow \overline{\bigcup_i \overline{E_i}} \in \mathcal{F} \xrightarrow{\text{De Morgan's Law}} \bigcap_i E_i \in \mathcal{F}.$$

□

In short, we can take unions, intersections, and complements of members of $\mathcal{F}$ in any combination and the result will always be a member of $\mathcal{F}$.

3.1.2 Probability Measure

(Kolmogorov's axioms of probability) Definition 4. A ***probability measure*** P is a function $P : \mathcal{F} \mapsto \mathbb{R}$ satisfying

(a) $P(E) \geq 0 \; \forall E \in \mathcal{F}$,

(b) $P(S) = 1$,

(c) If $E_1, E_2, \ldots \in \mathcal{F}$ are disjoint (i.e. $E_i \cap E_j = \emptyset \; \forall i \neq j$) then

$$P\left(\bigcup_{i=1}^{\infty} E_i\right) = \sum_{i=1}^{\infty} P(E_i).$$

The triplet $(S, \mathcal{F}, P)$, consisting of a set S, a σ-algebra $\mathcal{F}$ of subsets of S, and a probability measure P, is called a ***probability space***.

Comments

- The ***sample space*** (S) is the set of all possible outcomes of an experiment.
- The ***event space*** $(\mathcal{F})$ is the set of possible events, where an ***event*** E is a subset of the sample space, $E \subseteq S$. An ***elementary event*** is one that consist of a single element of S, i.e. a singleton.
- The probability measure (P) has three important interpretations:
  1. **classical**: Different outcomes in the sample space S are "equally likely",
  2. **frequentist**: the relative frequency of an event over many trials,
  3. **subjective**: a numerical measure of the degree of belief held by an individual.

  Example 5. "A sensor can detect items within 10 cm of the sensor. The sensor is placed in a room together with an object, and the probability that the sensor makes a detection is 0.0001."

  1. **classical**: The volume within 10 cm of the sensor divided by the volume of the room is 0.0001.
  2. **frequentist**: If we repeat the experiment a lot of times, then the fraction of the experiments in which the sensor makes a detection is 0.0001.
  3. **subjective**: Someone's subjective degree of belief, measured on a numerical scale from 0 to 1, that the sensor will detect is 0.0001.

- several results that can be derived from the probability measure axioms:
  - $P(\emptyset) = 0$.

  - $P(E) \leq 1$.
  - $P(\overline{E}) = 1 - P(E)$.
  - $P(E \cup F) = P(E) + P(F) - P(E \cap F)$..
  - $P(E \cap \overline{F}) = P(E) - P(E \cap F)$.
  - If $E \subset F$ then $P(E) \leq P(F)$.

3.2 Conditional Probability

Definition 6. If $P(F) > 0$ then the ***conditional probability*** of E given F is

$$P(E|F) = \frac{P(E \cap F)}{P(F)}.$$

Comments

- Difference among the following forms:
  - $P(E|F)$ – ***conditional probabilities***,
  - $P(E \cap F)$ – ***joint probabilities***,
  - $P(E)$ – ***marginal probabilities***.
- several results derived from the conditional probability definition:
  - $P(E|F) \geq 0$ for any event E.
  - $P(F|F) = 1$.
  - If the events $E_1, E_1, \ldots$ are pairwise disjoint, then $P\left(\left(\bigcup_i E_i\right)|F\right) = \sum_i P(E_i|F)$.
- Warning: In general, $P(E|F) \neq P(F|E)$.

Example 7. A medical test for a disease D has outcomes $+$ and $-$. The probabilities are

	D	$\overline{D}$	
$+$	0.009	0.099	0.108
$-$	0.001	0.891	0.892
	0.01	0.99	

By the definition of conditional probability, we have

$$P(+|D) = 90\%, \quad P(-|\overline{D}) = 90\%, \quad P(D|+) = \frac{0.009}{0.108} \approx 0.083.$$

The first two probabilities show that the test is fairly accurate. Sick people yield a positive 90% of the time and healthy people yield a negative 90% of the time.

3.3 Independence

Definition 8. Two events E and F are ***independent*** iff

$$P(E \cap F) = P(E)P(F).$$

Comments

- Extension: The events $E_1, \ldots, E_k$ are independent if, for every subset of events of size $l \leq k$, say indexed by $\{i_1, \ldots, i_l\}$,

$$P\left(\bigcap_{j=1}^{l} E_{i_j}\right) = \prod_{j=1}^{l} P(E_{i_j}).$$

- Independence could be either assumed or verified via the definition.
- Disjoint events with positive probability are not independent.
- From the definition of conditional probability, we can deduce that E and F are independent iff $P(E|F) = P(E)$.

Definition 9. For three events E_1, E_2, F, the pair of events E_1 and E_2 are said to be ***conditionally independent given*** F iff

$$P(E_1 \cap E_2|F) = P(E_1|F)P(E_2|F).$$

which could also be written as $E_1 \perp E_2 | F$.

3.4 Bayes' Theorem

(The Law of Total Probability) Theorem 10. Let $E_1, E_2, \ldots$ be a partition of S, i.e. $E_i \cap E_j = \emptyset$ for $i \neq j$ and $\bigcup_i E_i = S$. Then, for any event $F \subseteq S$, we have

$$P(F) = \sum_i P(F|E_i)P(E_i).$$

Proof. $P(F) = P(\bigcup_i F \cap E_i) = \sum_i P(F \cap E_i) = \sum_i P(F|E_i)P(E_i)$. □

(Bayes' Theorem) Theorem 11. If $P(F) > 0$ and let $E_1, E_2, \ldots$ be a partition on S s.t. $P(E_i) > 0\,\forall i$, we have

$$P(E_i|F) = \frac{P(F|E_i)P(E_i)}{P(F)} = \frac{P(F|E_i)P(E_i)}{\sum_j P(F|E_j)P(E_j)},$$

where $P(E_i|F)$ is called the ***posterior***, $P(F|E_i)$ is called the ***likelihood***, $P(E_i)$ is called the ***prior***, and $P(F)$ is called the ***evidence***.

Proof. Exercise! haha □

Example 12. A new covid-19 test is claimed to correctly identify 95% of people who are really covid-positive and 98% of people who are really covid-negative. If only 1 in a 1000 of the population are infected, what is the probability that a randomly selected person who tests positive actually has the disease?

Let I = "has a covid infection" and T = "test is positive". We are given $P(T|I) = 0.95, P(\overline{T}|\bar{I}) = 0.98, P(I) = 0.001$. We can thus derive that

$$P(I|T) = \frac{P(T|I)P(I)}{P(T)} = \frac{P(T|I)P(I)}{P(T|I)P(I) + P(T|\bar{I})P(\bar{I})} = \frac{0.95 \times 0.001}{0.95 \times 0.001 + 0.02 \times 0.999} = 0.045.$$

Chapter 4

Discrete Random Variables

4.1 Random Variables

Definition 13. A ***random variable*** is a (measurable) mapping

$$X : S \mapsto \mathbb{R}$$

with the property that $\{s \in S : X(s) \leq x\} \in \mathcal{F}\ \forall x \in \mathbb{R}$. This ensures that any set $B \subseteq \mathbb{R}$ corresponds to an event in the event space $\mathcal{F}$.

Definition 14. The image of S under X is called the ***range*** of the random variable

$$\mathbb{X} \equiv X(S) = \{X(s) | s \in S\} = \{x \in \mathbb{R} \,|\, \exists s \in S \text{ s.t. } X(s) = x\}.$$

So S contains all the possible outcomes of the experiment, $\mathbb{X}$ contains all the possible outcomes of the random variable X.

Definition 15. The ***probability distribution*** of X is defined as

$$P_X = P_X(X \in B \subseteq \mathbb{R}) = P(\{s \in S : X(s) \in B\})$$

which enables us to transfer the probability measure P defined on $\mathcal{F}$ to the real numbers in a natural way, and vice versa. For instance,

$$P_X(X = 7) = P(\{s \in S | X(s) = 7\}),$$
$$P_X(a < X \leq b) = P(\{s \in S | a < X(s) \leq b\}).$$

Example 16. Consider counting the number of heads in a sequence of 3 coin tosses. The underlying sample space is

$$S = \{TTT, TTH, THT, HTT, THH, HTH, HHT, HHH\}.$$

Since we are only interested in the number of heads in each sequence, we define the random variable X by

$$X(s) = \begin{cases} 0, & s = TTT, \\ 1, & s \in \{TTH, THT, HTT\}, \\ 2, & s \in \{HHT, HTH, THH\}, \\ 3, & s = HHH. \end{cases}$$

Thus, the probability of the number of heads X is less than 2 is

$$\begin{aligned} P_X(X < 2) &= P(\big\{s \in S : X(s) < 2\big\}) \\ &= P(\{TTT, TTH, THT, HTT\}) \\ &= \frac{\big|\{TTT, TTH, THT, HTT\}\big|}{|S|} \\ &= \frac{4}{8} = \frac{1}{2}. \end{aligned}$$

On a side note, the above process uses the classical interpretation on the probability measure.

Definition 17. The ***Cumulative Distribution Function (CDF)*** of a random variable X is the function $F_X : \mathbb{R} \mapsto [0, 1]$, defined by

$$F_X(x) = P_X(X \leq x) = P(\big\{s \in S : X(s) \leq x\big\}).$$

Comments

- Given a right-continuous function $F_X(x)$, check the following to verify if it is a valid CDF:
  (i) $0 \leq F_X(x) \leq 1 \; \forall x \in \mathbb{R}$,
  (ii) Monotonicity (non-decreasing): $\forall x_1, x_2 \in \mathbb{R}, x_1 < x_2 \Rightarrow F_X(x_1) \leq F_X(x_2)$.
  (iii) $F_X(-\infty) = 0, F_X(\infty) = 1$.
- For finite intervals $(a, b] \subseteq \mathbb{R}$, it is easy to check that
  $$P_X(a < X \leq B) = F_X(b) - F_X(a).$$
- Usually we suppress the subscript of $P_X(\cdot)$ and just write $P(\cdot)$ for the probability measure for the random variable, unless there is any ambiguity.

4.2 Discrete Random Variables

Definition 18. A random variable X is ***discrete*** if the range of X, $\mathbb{X}$, is countable, that is

$$\mathbb{X} = \{x_1, x_2, \ldots, x_n\} \text{ (finite)} \quad \text{or} \quad \mathbb{X} = \{x_1, x_2, \ldots\} \text{ (infinite)}.$$

Definition 19. For a discrete random variable X, we define the ***Probability Mass Function (PMF)*** as

$$p_X(x) = P_X(X = x), \quad x \in \mathbb{X}.$$

For completeness, we also define

$$p_X(x) = 0, \quad x \notin \mathbb{X}.$$

so that p_x is defined for all $x \in \mathbb{R}$.

Definition 20. The ***support*** of a random variable X is defined as

$$\left\{x \in \mathbb{R} : p_X(x) > 0\right\},$$

which is <u>almost always</u> the same as the range $\mathbb{X}$.

Properties of p_X and F_X

- $p_X(x_i) \geq 0$.
- $\sum_{x \in \mathbb{X}} p_X(x) = 1$.
- $F_X(x) = P(X \leq x)$, $x \in \mathbb{R}$.
- Let X be a discrete random variable with range $\mathbb{X} = \{x_1, x_2, \ldots\}$, where $x_1 < x_2 < \ldots$. Then for any $x \in \mathbb{R}$, if $x < x_1$, $F_X(x) = 0$; otherwise

  $$F_X(x) = \sum_{x_i \leq x} p_X(x_i) \iff p_X(x_i) = F_X(x_i) - F_X(x_{i-1}), \quad i = 2, 3, \ldots,$$

  with $p_X(x_1) = F_X(x_1)$.
- $\lim_{x \to -\infty} F_X(x) = 0$, $\lim_{x \to \infty} F_X(x) = 1$.
- F_X is continuous from the right on $\mathbb{R}$, i.e. for $x \in \mathbb{R}$, $\lim_{h \to 0^+} F_X(x + h) = F_X(x)$.
- F_X is non-decreasing, i.e. $a < b \Longrightarrow F_X(a) \leq F_X(b)$.
- For $a < b$, $P(a < X \leq b) = F_X(b) - F_X(a)$.

4.3 Functions of a Discrete Random Variable

Definition 21. The PMF of $Y = g(X)$ is found by grouping all the values in the range of x that correspond to the same value of Y, i.e.

$$p_Y(y) = \sum_{x \in \mathbb{X}: g(x)=y} p_X(x).$$

4.4 Mean and Variance

Definition 22. The ***expected value***, or ***mean*** of a discrete random variable X is defined to be

$$E_X(X) = \sum_{x \in \mathbb{X}} x p_X(x),$$

which is often written as $E(X)$, $E[X]$, or μ_X.

Theorem 23.

$$E(g(X)) = \sum_{x \in \mathbb{X}} g(x) p_X(x).$$

Proof. Let $Y = g(X)$, then

$$\begin{aligned} E(Y) &= \sum_{y \in \mathbb{Y}} y p_Y(y) \\ &= \sum_{y \in \mathbb{Y}} y \sum_{x \in \mathbb{X}: g(x)=y} p_X(x) \\ &= \sum_{y \in \mathbb{Y}} \sum_{x \in \mathbb{X}: g(x)=y} g(x) p_X(x) \\ &= \sum_{x \in \mathbb{X}} g(x) p_X(x). \end{aligned}$$

□

Theorem 24. Let X be a random variable with p_X. Let g and h be real-valued functions, $g, h : \mathbb{R} \mapsto \mathbb{R}$, and let a and b be constants. Then

$$E(ag(X) + bh(X)) = aE(g(X)) + bE(h(X)).$$

Proof. Exercise! □

Definition 25. Let X be a random variable. The ***variance*** of X, denoted by σ^2 or σ_X^2 or $\mathrm{Var}_X(X)$, is defined by

$$\mathrm{Var}_X(X) = E_X\left[(X - E_X(X))^2\right].$$

Proposition 26.

$$\mathrm{Var}(X) = E(X^2) - E(X)^2.$$

Proof.

$$\begin{aligned}\text{LHS} &= E\left[X^2 - 2E(X)X + E(X)^2\right] \\ &= E(X^2) - 2E(X)E(X) + E(X)^2 \\ &= \text{RHS}.\end{aligned}$$

□

Proposition 27.

$$\mathrm{Var}(aX \pm bY) = a^2\mathrm{Var}(X) + b^2\mathrm{Var}(Y) \pm 2ab\mathrm{Cov}(X, Y).$$

Proof. Exercise! □

Definition 28. The ***standard deviation*** of a random variable X, written $\mathrm{sd}_X(X)$ or σ_X, is the square root of the variance,

$$\sigma_X = \sqrt{\mathrm{Var}_X(X)}.$$

Definition 29. The ***skewness*** (γ_1) of a discrete random variable X is given by

$$\gamma_1 = \frac{E_X\left[\left\{X - E_X(X)\right\}^3\right]}{\sigma_X^3}.$$

Sums of Random Variables

Let $X_1, X_2, \ldots, X_n$ be n random variables, perhaps with different distributions and not necessarily independent. Let $S_n = \sum_{i=1}^n X_i$ be the sum of those variables, and $\frac{S_n}{n}$ be their sample average. Both S_n and $\overline{S} = \frac{S_n}{n}$ are random variables themselves.

The mean of S_n and $\frac{S_n}{n}$ are given by

$$E(S_n) = \sum_{i=1}^n E(X_i), \quad E\left(\overline{S}\right) = \frac{\sum_{i=1}^n E(X_i)}{n} = \mu_X.$$

If $X_1, X_2, \ldots, X_n$ are **independent**, we can calculate the variance of S_n and $\overline{S} = \frac{S_n}{n}$ as well:

$$\mathrm{Var}(S_n) = \sum_{i=1}^{n} \mathrm{Var}(X_i), \quad \mathrm{Var}(\overline{S}) = \frac{\sum_{i=1}^{n} \mathrm{Var}(X_i)}{n^2} = \frac{\sigma_X^2}{n}.$$

4.5 Some important Discrete Random Variables

Definition 30. We say X follows a ***Bernoulli Distribution*** if $X \sim \mathrm{Bernoulli}(p)$, where $0 \leq p \leq 1$, and the pmf is given by

$$p_X(x) = \begin{cases} p & x = 1 \\ 1-p & x = 0 \\ 0 & \text{otherwise} \end{cases} = p^x (1-p)^{1-x}, \quad x \in \mathbb{X} = \{0, 1\}.$$

Definition 31. We say X follows a ***Binomial Distribution*** if $X \sim \mathrm{Binomial}(n, p)$, where $0 \leq p \leq 1$ and $n \in \mathbb{Z}^+$, and the pmf is given by

$$p_X(x) = \binom{n}{x} p^x (1-p)^{n-x}, \quad x \in \mathbb{X} = \{0, 1, 2, \ldots, n\}.$$

Definition 32. We say X follows a ***Geometric Distribution*** if $X \sim \mathrm{Geometric}(p)$, where $0 \leq p \leq 1$, and the pmf is given by

$$p_X(x) = p(1-p)^{x-1}, \quad x \in \mathbb{X} = \{1, 2, \ldots\}.$$

Alternatively, let $Y = X - 1$, then $Y \sim \mathrm{Geometric}(p)$ with the pmf

$$p_Y(y) = p(1-p)^y, \quad y \in \mathbb{N} = \{0, 1, 2, \ldots\}.$$

Definition 33. We say X follows a ***Poisson Distribution*** if $X \sim \mathrm{Poissons}(\lambda)$, where $\lambda > 0$, and the pmf is given by

$$p_X(x) = \frac{e^{-\lambda} \lambda^x}{x!}, \quad x \in \mathbb{X} = \{0, 1, 2, \ldots\}.$$

Definition 34. We say X follows a ***Discrete Uniform Distribution*** if $X \sim$ Uniform({1, 2, . . . , n}), and the pmf is given by

$$p_X(x) = \frac{1}{n}, \quad x \in \mathbb{X} = \{1, 2, \ldots, n\}.$$

Table 4.1: Means and Variances of different distributions

	Mean(μ)	Variance(σ^2)	Skewness(γ_1)
Bernoulli	p	$p(1-p)$	N.A.
Binomial	np	$np(1-p)$	$\frac{1-2p}{\sqrt{np(1-p)}}$
Geometric(original)	$\frac{1}{p}$	$\frac{1-p}{p^2}$	$\frac{2-p}{\sqrt{1-p}}$
Geometric(alternative)	$\frac{1-p}{p}$	$\frac{1-p}{p^2}$	$\frac{2-p}{\sqrt{1-p}}$
Poisson	λ	λ	$\frac{1}{\sqrt{\lambda}}$
Uniform	$\frac{n+1}{2}$	$\frac{n^2-1}{12}$	0

Comments

- From table 4.1, we can see that the skewness of both Geometric and Poisson Distribution is always positive.
- **Approximation of Bionomial distribution as Poisson distribution**. It can be shown that for Binomial(n, p), <u>when p is small and n is large</u>, this distribution can be well approximated by the Poisson distribution with rate parameter $\lambda = np$, Poisson(np).

Chapter 5

Continuous Random Variable

5.1 Definitions

Definition 35. A random variable X is (absolutely) continuous if $\exists\, f_X : \mathbb{R} \mapsto \mathbb{R}$ (measurable) s.t. f_X is non-negative and

$$P(X \in B) = \int_{x \in B} f_X(x)\mathrm{d}x, \quad B \subseteq \mathbb{R},$$

and f_X is referred to as the ***Probability Density Function (PDF)*** of X.

Comments

- It follows that f_X is a pdf for a continuous variable X iff

$$f_X(x) \geq 0 \quad \text{and} \quad \int_{-\infty}^{\infty} f_X(x)\mathrm{d}x = 1.$$

- The pdf $f_X(x)$ is *not* a probability. It is a probability *density*, having units of 1/[units of X]. As such,

$$\forall x \in \mathbb{R}, P(X = x) = 0.$$

- Since the pdf is not itself a probability, unlike the pmf of a discrete random variable, we do *not* require $f_X(x) \leq 1$.

Definition 36. The ***Cumulative Distribution Function (CDF)***, F_X, of a continuous random variable X is defined as

$$F_X(x) = P(X \leq x), x \in \mathbb{R}.$$

Comments

- From now on, we implicitly assume the absolutely continuous case, then the CDF can be written as
$$F_X(x) = P(X \le x) = \int_{-\infty}^{x} f_X(x')\mathrm{d}x', \quad x \in \mathbb{R}.$$
- For the cdf of a continuous random variable,
$$\lim_{x \to -\infty} F_X(x) = 0, \quad \lim_{x \to \infty} F_X(x) = 1.$$
- At values of x where F_X is differentiable,
$$f_X(x) = \frac{\mathrm{d}}{\mathrm{d}t} F_X(t)\bigg|_{t=x} \equiv F_X'(x).$$
- For $a < b$,
$$P(a < X \le b) = P(a \le X < b) = P(a \le X \le b) = P(a < X < b) = F_X(b) - F_X(a).$$

5.2 Transformations

Let X be a continuous random variable with pdf f_X and cdf F_X. Let $Y = g(X)$ be a *transformation* (function) of X for some (measurable) function $g : \mathbb{R} \mapsto \mathbb{R}$ s.t. g is continuous. Given f_X, how do we obtain f_Y?

Method 1

1. Integrate $f_X(x)$ to find $F_X(x)$.
2. Find $F_Y(y)$ in terms of $F_X(x)$.
3. Differentiate $F_Y(y)$ to get pdf $f_Y(y)$.

Example 37. Given $f_X(x) = e^{-x}$ for $x > 0$. Thus,
$$F_X(x) = \int_0^x f_X(u)\mathrm{d}u = 1 - e^{-x}.$$
Let $Y = g(X) = \log(X)$. Then the range of Y is $\mathbb{R}$ and
$$F_Y(y) = P(Y \le y) = P(\log(X) \le y) = P(X \le e^y) = F_X(e^y).$$
Taking the derivative of the cdf gives the pdf
$$f_Y(y) = \frac{\mathrm{d}}{\mathrm{d}y} F_Y(y) = \frac{\mathrm{d}}{\mathrm{d}y} F_X(e^y) = e^y f_X(e^y) = e^y e^{-e^y}.$$

Method 2

1. Go to the pdf directly by matching the equation $f_Y(y)\mathrm{d}y = f_X(x)\mathrm{d}x$.

Example 38. Given $f_X(x) = e^{-x}$ for $x > 0$ and let $Y = g(X) = \log(X)$. Then the range of Y is $\mathbb{R}$. We can then obtain

$$x = g^{-1}(y) = e^y, \quad \frac{\mathrm{d}y}{\mathrm{d}x} = \frac{1}{x} \Rightarrow \mathrm{d}x = |x\mathrm{d}y| = e^y\mathrm{d}y.$$

The absolute sign is to ensure that the product $f_X(x_i)\mathrm{d}x_i$ is not negative. Fitting into the equation, we have

$$f_Y(y)\mathrm{d}y = f_X(x)\mathrm{d}x = f_X(e^y)e^y\mathrm{d}y.$$

We can thus obtain

$$f_Y(y) = f_X(e^y)e^y = e^y e^{-e^y}.$$

Warning g may not be a 1-to-1 function, e.g. $Y = X^2$. In this case, always draw a graph and think about the ranges of X and Y. Following the example of $Y = X^2$, we can derive that

$$x = \pm\sqrt{y}, \quad \frac{\mathrm{d}y}{\mathrm{d}x} = 2x = \pm 2\sqrt{y},$$

and then note the following

$$f_Y(y)\mathrm{d}y = f_X(x)\mathrm{d}x = f_X(\sqrt{y})\left|\frac{\mathrm{d}y}{2\sqrt{y}}\right| + f_X(-\sqrt{y})\left|\frac{\mathrm{d}y}{-2\sqrt{y}}\right|,$$

$$\Rightarrow f_Y(y) = \frac{f_X(\sqrt{y}) + f_X(-\sqrt{y})}{2\sqrt{y}}.$$

Example 39. $X \sim \text{Uniform}(-1, 3)$, let $Y = X^2$. Find $f_Y(y)$.

Firstly, we have

$$f_X(x) = \begin{cases} \frac{1}{4}, & -1 \le x \le 3, \\ 0, & \text{otherwise} \end{cases}, \quad f_Y(y) = \frac{\mathrm{d}x}{\mathrm{d}y} f_X(x).$$

And we can also obtain that

$$x = \pm\sqrt{y}, \quad \frac{\mathrm{d}x}{\mathrm{d}y} = \pm\frac{1}{2\sqrt{y}}.$$

Thus when $0 \le y \le 1$:

$$f_Y(y) = \frac{1}{2\sqrt{y}}\left(\frac{1}{4} + \frac{1}{4}\right) = \frac{1}{4\sqrt{y}}.$$

and when $1 < y \le 9$:

$$f_Y(y) = \frac{1}{2\sqrt{y}}\left(\frac{1}{4}\right) = \frac{1}{8\sqrt{y}}.$$

Finally for other values of y, we have $f_Y(y) = 0$.

5.3 Mean, Variance and Quantiles

Definition 40. For a continuous random variable X we define the ***mean*** or ***expectation*** of X as

$$\mu_X \text{ or } E_X(X) = \int_{-\infty}^{\infty} x f_X(x)\mathrm{d}x.$$

Comments

- More generally, for a (measurable) function of interest $g : \mathbb{R} \mapsto \mathbb{R}$ we have

$$E_X(g(X)) = \int_{-\infty}^{\infty} g(x) f_X(x)\mathrm{d}x.$$

- Linearity of expectation:

$$E[ag(X) + b] = aE[g(X)] + b, \quad \forall a, b \in \mathbb{R}, g : \mathbb{R} \mapsto \mathbb{R}.$$

Definition 41. The ***variance*** of a continuous random variable X is given by

$$\sigma_X^2 \text{ or } \mathrm{Var}_X(X) = E[(X - \mu_X)^2] = \int_{-\infty}^{\infty} (x - \mu_X)^2 f_X(x)\mathrm{d}x.$$

Comments

- Equivalently,

$$\mathrm{Var}_X(X) = \int_{-\infty}^{\infty} x^2 f_X(x)\mathrm{d}x - \mu_X^2 = E(X^2) - E(X)^2.$$

- For a linear transformation $aX + b$ we again have

$$\mathrm{Var}(aX + b) = a^2 \mathrm{Var}(X), \quad \forall a, b \in \mathbb{R}.$$

Definition 42. For a (continuous) random variable X we define the α***-quantile*** $Q_X(\alpha)$, $0 \leq \alpha \leq 1$ to satisfy $P(X \leq Q_X(\alpha)) = \alpha$,

$$Q_X(\alpha) = F_X^{-1}(\alpha).$$

In particular, the ***median*** of X is $Q_X\left(\frac{1}{2}\right)$.

5.4 Some Important Continuous Random Variables

Definition 43. We say that X follows a ***continuous uniform distribution*** on the interval (a, b), where $a < b$, if $X \sim \mathrm{U}(a, b)$, and the pdf is given by

$$f_X(x) = \begin{cases} \dfrac{1}{b-a}, & a < x < b, \\ 0, & \text{otherwise.} \end{cases}$$

The cdf is given by

$$F_X(x) = \begin{cases} 0, & x \leq a, \\ \dfrac{x-a}{b-a}, & a < x < b, \\ 1, & x \geq b. \end{cases}$$

The distribution $X \sim \mathrm{U}(0, 1)$ is referred to as the ***standard uniform distribution***.

Definition 44. We say that X follows a ***exponential distribution*** if $X \sim \mathrm{Exp}(\lambda)$, where $\lambda > 0$, and the pdf is given by

$$f_X(x) =^{-\lambda x}, \quad x \geq 0.$$

The cdf is given by

$$F_X(x) = 1 - e^{-\lambda x}, \quad x \geq 0.$$

Comments

- Interpretation: For $T \sim \mathrm{Exp}(\lambda)$, T can be interpreted as the time until an event occurs, where events occur at an "average rate" λ. The exponential distribution is the continuous version of the geometric distribution.

- "Lack of memory": If we have already waited for a time t, what is the probability of still waiting at time $t + s$?

$$\begin{aligned} P(X > s + t \,|\, X > t) &= \frac{P(X > s + t \cap X > t)}{P(X > t)} \\ &= \frac{P(X > s + t)}{P(X > t)} \\ &= \frac{e^{-\lambda(s+t)}}{e^{-\lambda t}} = e^{-\lambda s} \\ &= P(X > s). \end{aligned}$$

  In words, the knowledge that we have waited for time s for an event tells us nothing about how much longer we will have to wait, i.e. the process has *no memory*. This is known as the **Lack of Memory** property, and is unique to the exponential distribution amongst continuous distributions.

- Relation with Poisson distribution: If events in a random process occur according to a Poisson distribution with rate λ, then the time between events has an Exponential distribution with rate parameter λ.

  Proof. Suppose we have some random event process such that $\forall x > 0$, the number of events occuring in $[0, x]$, N_x, follows a Poisson distribution with rate parameter λ, so $N_x \sim \text{Poisson}(\lambda x)$. Such a process is known as a *Homogeneous Poisson process*. Let X be the time until the first event of this process arrives. Then we notice that

  $$P(X > x) = P(N_x = 0) = \frac{(\lambda x)^0 e^{-\lambda x}}{0!} = e^{-\lambda x}.$$

  Hence $X \sim \text{Exp}(\lambda)$. This argument applies for all subsequent inter-arrival times. $\square$

Definition 45. We say that X follows a ***Gaussian*** or ***normal distribution*** if $X \sim \text{N}(\mu, \sigma^2)$, where $\mu \in \mathbb{R}, \sigma > 0$, and the pdf is given by

$$f_X(x) = \frac{1}{\sigma\sqrt{2\pi}} \exp\left\{-\frac{(x-\mu)^2}{2\sigma^2}\right\}.$$

The cdf is given by

$$F_X(x) = \frac{1}{\sigma\sqrt{2\pi}} \int_{-\infty}^{x} \exp\left\{-\frac{(t-\mu)^2}{2\sigma^2}\right\} \mathrm{d}t.$$

Comments

- If $X \sim \text{N}(0,1)$, then X has a ***standard*** or ***unit normal distribution***. The pdf of the standard normal distribution is written as $\phi(x)$, and the cdf is written as $\Phi(x)$.
- If $Y \sim \text{N}(0,1)$, and $X = \sigma Y + \mu$, (or equivalently $Y = \frac{X-\mu}{\sigma}$) then $X \sim \text{N}(\mu, \sigma^2)$. We can then write the cdf of X in terms of Φ,

  $$F_X(x) = P(X \le x) = P\left(Y \le \frac{x-\mu}{\sigma}\right) = \Phi\left(\frac{x-\mu}{\sigma}\right).$$

- Because the standard normal pdf ϕ is *symmetric* about 0, i.e. $\phi(-z) = \phi(z)$ for $z \in \mathbb{R}$, for the cdf we have

  $$\Phi(z) = 1 - \Phi(-z).$$

Table 5.1: Means and Variances of different continuous distributions

	Mean(μ)	Variance(σ^2)
Uniform	$\frac{a+b}{2}$	$\frac{(b-a)^2}{12}$
Exponential	$\frac{1}{\lambda}$	$\frac{1}{\lambda^2}$
Normal	μ	σ

Chapter 6

Jointly Distributed Random Variables

6.1 Definitions

Definition 46. Given a pair of random variables, X and Y, we define the ***joint probability distribution*** P_{XY} as follows:

$$\begin{aligned} P_{XY}(B_X, B_Y) &= P\left(X^{-1}(B_X) \cap Y^{-1}(B_Y)\right) \\ &= P\left(\left\{s \in S : X(s) \in B_X, Y(s) \in B_Y\right\}\right), \quad B_X, B_Y \subseteq \mathbb{R}. \end{aligned}$$

More generally, for some $B_{XY} \subseteq \mathbb{R}^2$, find the collection of sample space elements (i.e. the event)

$$S_{XY} = \left\{s \in S : \begin{pmatrix} X(s) & Y(s) \end{pmatrix} \in B_{XY}\right\},$$

and define

$$P_{XY}(B_{XY}) = P(S_{XY}).$$

Definition 47. Given a pair of random variables, X and Y, the ***joint cumulative distribution function*** is defined as

$$F_{XY}(x, y) = P_{XY}(X \leq x, Y \leq y), \quad x, y \in \mathbb{R}.$$

Comments

- The marginal cdfs of X and Y can be recovered by

$$F_X(x) = F_{XY}(x, \infty), \quad F_Y(y) = F_{XY}(\infty, y), \quad x, y \in \mathbb{R}.$$

- $\forall x, y \in \mathbb{R}$,

$$0 \leq F_{XY}(x, y) \leq 1,$$

$$F_{XY}(x, -\infty) = 0, \quad F_{XY}(-\infty, y) = 0, \quad F_{XY}(\infty, \infty) = 1.$$

- Monotonicity: $\forall x, y \in \mathbb{R}$, we have

$$x_1 < x_2 \Rightarrow F_{XY}(x_1, y) \leq F_{XY}(x_2, y), \quad y_1 < y_2 \Rightarrow F_{XY}(x, y_1) \leq F_{XY}(x, y_2).$$

- By noting that $P_{XY}(x_1 < X \leq x_2, Y \leq y) = F_{XY}(x_2, y) - F_{XY}(x_1, y)$, we can also obtain that

$$P_{XY}(x_1 < X \leq x_2, y_1 < Y \leq y_2) = F_{XY}(x_2, y_2) - F_{XY}(x_1, y_2) - F_{XY}(x_2, y_1) + F_{XY}(x_1, y_1).$$

Definition 48. If X and Y are both discrete random variables, then we can define the ***joint probability mass function*** as

$$p_{XY}(x, y) = P_{XY}(X = x, Y = y), \quad x, y \in \mathbb{R}.$$

Comments

- We can recover the marginal pmfs p_X and p_Y, by the law of total probability, $\forall x, y \in \mathbb{R}$,

$$p_X(x) = \sum_{y \in \mathbb{Y}} p_{XY}(x, y), \quad p_Y(y) = \sum_{x \in \mathbb{X}} p_{XY}(x, y).$$

- For p_{XY} to be a valid pmf, we need to make sure the following conditions hold:

$$0 \leq p_{XY}(x, y) \leq 1, \forall x, y \in \mathbb{R} \quad \text{and} \quad \sum_{y \in \mathbb{Y}} \sum_{x \in \mathbb{X}} p_{XY}(x, y) = 1.$$

Definition 49. If $\exists f_{XY} : \mathbb{R} \times \mathbb{R} \mapsto \mathbb{R}$ s.t. $\forall B_{XY} \subseteq \mathbb{R} \times \mathbb{R}$,

$$P_{XY}(B_{XY}) = \int_{(x,y) \in B_{XY}} f_{XY}(x, y) \mathrm{d}x \mathrm{d}y,$$

then we say X and Y are ***jointly continuous*** and we refer to f_{XY} as the ***joint probability density function*** of X and Y. In this case we have

$$F_{XY}(x, y) = \int_{t=-\infty}^{y} \int_{s=-\infty}^{x} f_{XY}(s, t) \mathrm{d}s \mathrm{d}t, \quad x, y \in \mathbb{R}.$$

By the fundamental theorem of calculus we can identify the joint pdf of X and Y as

$$f_{XY}(x, y) = \frac{\partial^2}{\partial x \partial y} F_{XY}(x, y).$$

Comments

- We can recover the marginal densities f_X as

$$f_X(x) = \frac{\mathrm{d}}{\mathrm{d}x} F_X(x) = \frac{\mathrm{d}}{\mathrm{d}x} F_{XY}(x, \infty) = \frac{\mathrm{d}}{\mathrm{d}x} \int_{y=-\infty}^{\infty} \int_{s=-\infty}^{x} f_{XY}(s, y) \mathrm{d}s \mathrm{d}y,$$

  and by the fundamental theorem of calculus, we obtain (similarly for f_Y)

$$f_X(x) = \int_{y=-\infty}^{\infty} f_{XY}(x, y) \mathrm{d}y, \quad f_Y(y) = \int_{x=-\infty}^{\infty} f_{XY}(x, y) \mathrm{d}x.$$

- For f_{XY} to be a valid pdf, we need to make sure the following conditions hold:

$$f_{XY}(x, y) \geq 0, \forall x, y \in \mathbb{R} \quad \text{and} \quad \int_{y=-\infty}^{\infty} \int_{x=-\infty}^{\infty} f_{XY}(x, y) \mathrm{d}x \mathrm{d}y = 1.$$

Example 50. Suppose continuous random variables $(X, Y) \in \mathbb{R}^2$ have joint pdf

$$f(x, y) = \begin{cases} 1, & |x| + |y| < \frac{1}{\sqrt{2}} \\ 0, & \text{otherwise.} \end{cases}$$

Determine the marginal pdfs of X and Y.

$$f_X(x) = \int_{-\infty}^{\infty} f(x, y) \mathrm{d}y = \int_{|x| - \frac{1}{\sqrt{2}}}^{\frac{1}{\sqrt{2}} - |x|} \mathrm{d}y = \sqrt{2} - 2|x|.$$

Similarly, we can obtain $f_Y(y) = \sqrt{2} - 2|y|$.

6.2 Independence, Conditional Probability, Expectation

Definition 51. Two continuous random variables X and Y are ***independent*** iff

$$f_{XY}(x, y) = f_X(x) f_Y(y), \quad \forall x, y \in \mathbb{R}.$$

Example 52. Suppose that the lifetime, X, and brightness, Y, of a light bulb are modelled as continuous random variables. Let their joint pdf be given by

$$f(x, y) = \lambda_1 \lambda_2 e^{-\lambda_1 x - \lambda_2 y}, \quad x, y > 0.$$

Are lifetime and brightness independent?

$$f_X(x) = \int_0^{\infty} \lambda_1 \lambda_2 e^{-\lambda_1 x - \lambda_2 y} \mathrm{d}y$$

$$\begin{aligned}
&= \lambda_1 e^{-\lambda_1 x} \int_0^\infty \lambda_2 e^{-\lambda_2 y} \mathrm{d}y \\
&= \lambda_1 e^{-\lambda_1 x} \left[-e^{-\lambda_2 y}\right]_0^\infty \\
&= \lambda_1 e^{-\lambda_1 x}.
\end{aligned}$$

Similarly, we have $f_Y(y) = \lambda_2 e^{-\lambda_2 y}$. Thus we obtain that $f_X(x) f_Y(y) = f_{XY}(x,y)$, indicating that the lifetime and brightness are independent.

Definition 53. For two random variables X and Y, we define the ***conditional probability distribution*** $P_{Y|X}$ by

$$P_{Y|X}(B_Y|B_X) = \frac{P_{XY}(B_X, B_Y)}{P_X(B_X)}, \quad B_X, B_Y \subseteq \mathbb{R}.$$

Definition 54. For random variables X and Y, we define the ***conditional probability density function*** $f_{Y|X}$ by

$$f_{Y|X}(y|x) = \frac{f_{XY}(x,y)}{f_X(x)}, \quad x, y \in \mathbb{R}.$$

Comments

The random variables X and Y are independent

$$\begin{aligned}
&\iff P_{Y|X}(B_Y|B_X) = P_Y(B_Y), \qquad && \forall B_X, B_Y \subseteq \mathbb{R}, \\
&\iff f_{Y|X}(y|x) = f_Y(y), && \forall x, y \in \mathbb{R}.
\end{aligned}$$

Definition 55. If X and Y are discrete, we define $E(g(X,Y))$ by

$$E(g(X,Y)) = \sum_{y \in \mathbb{Y}} \sum_{x \in \mathbb{X}} g(x,y) p_{XY}(x,y).$$

If X and Y are jointly continuous, we define $E(g(X,Y))$ by

$$E(g(X,Y)) = \int_{y=-\infty}^{\infty} \int_{x=-\infty}^{\infty} g(x,y) f_{XY}(x,y) \mathrm{d}x \mathrm{d}y.$$

Comments

- Expectation is always linear:

$$E_{XY}[g_1(X,Y) + g_2(X,Y)] = E_{XY}[g_1(X,Y)] + E_{XY}[g_2(X,Y)].$$

- If $g(X, Y) = g_1(X)g_2(Y)$ and X and Y are **independent**,
$$E_{XY}[g_1(X)g_2(Y)] = E_X[g_1(X)]E_Y[g_2(Y)].$$
In particular, considering $g(X, Y) = XY$ for independent X and Y,
$$E_{XY}(XY) = E_X(X)E_Y(Y).$$
Warning! In general $E_{XY}(XY) \neq E_X(X)E_Y(Y)$.

Definition 56. If X and Y are discrete, the ***conditional expectation*** of Y given $X = x$ is
$$E_{Y|X}(Y|X = x) = \sum_{y \in \mathbb{Y}} y\, p(y|x).$$
Similarly for the case when X and Y are continuous, we have
$$E_{Y|X}(Y|X = x) = \int_{y=-\infty}^{\infty} y\, f(y|x)\mathrm{d}y.$$

Definition 57. We define the ***correlation*** of X and Y by
$$\rho_{XY} = \mathrm{Cor}(X, Y) = \frac{\sigma_{XY}}{\sigma_X \sigma_Y}.$$

6.3 Multivariate Transformations

(Convolution Theorem) Theorem 58. If X and Y are *independent* random variables and $Z = X + Y$, then
$$p_Z(z) \text{ or } f_Z(z) = \begin{cases} \displaystyle\sum_{x \in \mathbb{X}} p_X(x)p_Y(z - x) & \text{(discrete case)}, \\ \displaystyle\int_{\mathbb{R}} f_X(\omega)f_Y(z - \omega)\mathrm{d}\omega & \text{(continuous case)}. \end{cases}$$

Example 59. Supposed $X \sim \mathrm{N}(0, \sigma^2)$ and $Y \sim \mathrm{N}(0, 1)$ with X and Y independent. Let $Z = X + Y$ and derive the pdf of Z.
$$\begin{aligned} f_Z(z) &= \int_{-\infty}^{\infty} f_X(x)f_Y(z - x)\mathrm{d}x \\ &= \int_{-\infty}^{\infty} \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{x^2}{2\sigma^2}} \frac{1}{\sqrt{2\pi}} e^{-\frac{(z-x)^2}{2}} \mathrm{d}x \\ &\vdots \\ &= \frac{1}{\sqrt{2\pi(1 + \sigma^2)}} e^{-\frac{z^2}{2(1+\sigma^2)}}. \\ &\implies Z \sim \mathrm{N}(0, 1 + \sigma^2). \end{aligned}$$

Theorem 60. If $X \sim \mathrm{N}(\mu_X, \sigma_X^2)$ and $Y \sim \mathrm{N}(\mu_Y, \sigma_Y^2)$ with X and Y independent, then

$$Z = X + Y \sim \mathrm{N}(\mu_X + \mu_Y, \sigma_X^2 + \sigma_Y^2).$$

Theorem 61. Suppose (X, Y) is a bivariate random variable and let $U = g_1(X, Y)$ and $V = g_2(X, Y)$. Then for any $B \subseteq \mathbb{R}^2$,

$$P\big((U, V) \in B\big) = P\big((X, Y) \in A\big),$$

where $A = \big\{(x, y) : (g_1(x, y), g_2(x, y)) \in B\big\}$. This can be generally divided into two cases to consider:

1. If (X, Y) is *discrete*: Let $A(u, v) = \big\{(x, y) \in (\mathbb{X}, \mathbb{Y}) : (g_1(x, y), g_2(x, y)) = (u, v)\big\}$, then

$$p_{UV}(u, v) = P(U = u, V = v) = P\big((X, Y) \in A(u, v)\big) = \sum_{\substack{(x,y): \\ g_1(x,y)=u, \\ g_2(x,y)=v}} p_{XY}(x, y).$$

2. If (X, Y) is *continuous*: We define the ***Jacobian determinant*** $|J|$ s.t. $\mathrm{d}x\mathrm{d}y = |J|\mathrm{d}u\mathrm{d}v$, where

$$|J| = \left|\frac{\partial(x, y)}{\partial(u, v)}\right| = \left|\det\begin{pmatrix} \dfrac{\partial x}{\partial u} & \dfrac{\partial x}{\partial v} \\ \dfrac{\partial y}{\partial u} & \dfrac{\partial y}{\partial v} \end{pmatrix}\right| = \left|\frac{\partial x}{\partial u}\frac{\partial y}{\partial v} - \frac{\partial x}{\partial v}\frac{\partial y}{\partial u}\right|.$$

Then

$$f_{UV}(u, v) = f_{XY}(x, y)\left|\frac{\partial(x, y)}{\partial(u, v)}\right|.$$

6.4 Gamma and Beta Distributions

Definition 62. The ***Gamma function*** is defined by

$$\Gamma(\alpha) = \int_0^\infty t^{\alpha-1}e^{-t}\mathrm{d}t, \quad \alpha > 0,$$

and then we say X follows the ***Gamma Distribution*** if $X \sim \mathrm{Gamma}(\alpha, \beta)$, where $\alpha, \beta > 0$, and we have

$$f_X(x) = \frac{\beta^\alpha}{\Gamma(\alpha)}x^{\alpha-1}e^{-x\beta}, \quad x \in (0, \infty).$$

Comments

- $\Gamma(\alpha) = (\alpha - 1)\Gamma(\alpha - 1)$ for $\alpha > 1$.
- $\Gamma(1) = \int_0^\infty e^{-t}\mathrm{d}t = 1$.
- $\Gamma(n) = (n-1)!$ for $n \in \mathbb{Z}^+$.
- $\Gamma(\frac{1}{2}) = \sqrt{\pi}$.

Theorem 63. If $X \sim \text{Gamma}(\lambda, \theta)$ and $Y \sim \text{Gamma}(\xi, \theta)$ with X and Y independent, then $Z = X + Y \sim \text{Gamma}(\lambda + \xi, \theta)$.

Definition 64. The ***Beta function*** is defined by

$$B(\alpha, \beta) = \int_0^1 x^{\alpha-1}(1-x)^{\beta-1}\mathrm{d}x = \frac{\Gamma(\alpha)\Gamma(\beta)}{\Gamma(\alpha+\beta)}.$$

We say X follows the ***Beta Distribution*** if $X \sim \text{Beta}(\alpha, \beta)$, where $\alpha, \beta > 0$, and we have

$$f_X(x) = \frac{x^{\alpha-1}(1-x)^{\beta-1}}{B(\alpha, \beta)}, \quad 0 < x < 1,$$

Chapter 7

Convergence Concepts and Theorems

7.1 Modes of Convergence

Definition 65. Let $X_1, X_2, \ldots, X_n, X$ be random variables. Higher order of strength implies the lower ones. In decreasing order of strength, we have

1. X_n ***converges almost surely*** to X if
$$P(\lim_{n\to\infty} X_n = X) = 1 \quad \text{or} \quad X_n \to_{\text{as}} X.$$
2. X_n ***converges in probability*** to X if
$$\forall \epsilon > 0, \lim_{n\to\infty} P(|X_n - X| > \epsilon) = 0 \quad \text{or} \quad X_n \to_{\text{p}} X.$$
3. X_n ***converges in distribution (converges weakly)*** to X with cdf F_X if
$$\lim_{n\to\infty} P(X_n < x) = F_X(x) \quad \text{or} \quad X_n \to_{\text{d}} X.$$
at all continuity points x of $F_X(x)$.
4. If $X_n \to_{\text{d}} X$ and $P(X = c) = 1$ for some c, we say the limiting distribution of X_n is ***degenerate*** at c and write $X_n \to_{\text{d}} c$.

Comments

- *Convergence in distribution* only requires the cdf of X_n's converges to the cdf of X as $n \to \infty$. It does not require any dependence between the X_n's and X, i.e. it doesn't tell anything about whether the value of X_n will be close to X for a single run of the experiment. Thus it is in some sense the weakest type of convergence. Convergence in probability says that for large enough n, *for each run of the experiment*, there is a high probability that the two values, X_n and X, will be close together.

- $X_n \to_{\text{d}} c \iff X_n \to_{\text{p}} c$ for some c.

- (Slutsky's Theorem) If $X_n \to_{\mathrm{d}} X$ and $Y_n \to_{\mathrm{d}} c$, then

$$X_n Y_n \to_{\mathrm{d}} cX \quad \text{and} \quad X_n + Y_n \to_{\mathrm{d}} X + c.$$

Example 66. Let $X_1, X_2, X_3, \ldots$ be a sequence of random variables s.t.

$$F_{X_n}(x) = \begin{cases} 1 - (1 - \frac{1}{n})^{nx} & x > 0 \\ 0 & \text{otherwise.} \end{cases}$$

Show that $X_n \to_{\mathrm{d}}$ Exponential(1).

Let $X \sim$ Exponential(1). For $x \leq 0$, we have

$$F_{X_n}(x) = F_X(x) = 0, \quad \text{for } n = 1, 2, 3, \ldots.$$

For $x > 0$, we have

$$\begin{aligned} \lim_{n\to\infty} F_{X_n}(x) &= \lim_{n\to\infty} \left(1 - \left(1 - \frac{1}{n}\right)^{nx}\right) \\ &= 1 - \lim_{n\to\infty} \left(1 - \frac{1}{n}\right)^{nx} \\ &= 1 - e^{-x} \\ &= F_X(x). \end{aligned}$$

Thus, we conclude that $X_n \to_{\mathrm{d}} X$.

Example 67. Let X be a random variable, and $X_n = X + Y_n$, where

$$E(Y_n) = \frac{1}{n} \quad \text{and} \quad (Y_n) = \frac{\sigma^2}{n},$$

where $\sigma > 0$ is a constant. Show that $X_n \to_{\mathrm{p}} X$.

By the triangle inequality, $\forall a, b \in \mathbb{R}, |a + b| \leq |a| + |b|$. Choosing $a = Y_n - E(Y_n)$ and $b = E(Y_n)$, we obtain

$$|Y_n| \leq |Y_n - E(Y_n)| + \frac{1}{n}.$$

Now for any $\epsilon > 0$, we have

$$\begin{aligned} P(|X_n - X| \geq \epsilon) &= P(|Y_n| \geq \epsilon) \\ &\leq P(|Y_n - E(Y_n)| + \frac{1}{n} \geq \epsilon) \\ &= P(|Y_n - E(Y_n)| \geq \epsilon - \frac{1}{n}) \\ &\leq \frac{(Y_n)}{\left(\epsilon - \frac{1}{n}\right)^2} \quad \text{(by Chebyshev's inequality)} \\ &= \frac{\sigma^2}{n\left(\epsilon - \frac{1}{n}\right)^2} \to 0 \quad \text{as } n \to \infty. \end{aligned}$$

Therefore, we conclude that $X_n \to_{\mathrm{p}} X$.

7.2 Moment Generating Functions

Definition 68. The ***moment generating function (MGF)*** of the random variable X is

$$M_X(t) = E\left(e^{tX}\right),$$

provided the expectation exists in some neighborhood of zero, i.e. the expectation exists $\forall |t| < \epsilon$ for some $\epsilon > 0$.

Theorem 69. If X has an MGF, then

$$E(X^n) = M_X^{(n)}(0) = \frac{\mathrm{d}^n}{\mathrm{d}t^n} M_X(t)\bigg|_{t=0}.$$

Proof. When $n = 1$,

$$\frac{\mathrm{d}}{\mathrm{d}t} M_X(t) = \frac{\mathrm{d}}{\mathrm{d}t} \int_{-\infty}^{\infty} e^{tx} f_X(x) \mathrm{d}x = \int_{-\infty}^{\infty} x e^{tx} f_X(x) \mathrm{d}x = E\left[X e^{tX}\right].$$

When $t = 0$, $M_X^{(1)}(0) = E(X)$. It is similar for $n > 1$. Or we could also observe the different moments using Taylor expansions at $t = 0$, that

$$e^{tX} = 1 + tX + \frac{(tX)^2}{2!} + \frac{(tX)^3}{3!} + \dots$$

leading to

$$E\left[e^{tX}\right] = 1 + t \underbrace{E(X)}_{1^{\text{st}} \text{ moment}} + \frac{t^2}{2!} \underbrace{E(X^2)}_{2^{\text{nd}} \text{ moment}} + \frac{t^3}{3!} \underbrace{E(X^3)}_{3^{\text{rd}} \text{ moment}} + \dots.$$

$\square$

Example 70. Suppose $X \sim \mathrm{N}(0, 1)$. Derive the MGF and the first four moments of X.

$$\begin{aligned}
M_X(t) &= E(e^{tX}) = \int_{-\infty}^{\infty} e^{tx} \frac{1}{\sqrt{2\pi}} e^{-\frac{1}{2}x^2} \mathrm{d}x = e^{\frac{1}{2}t^2}, \\
M_X^{(1)}(t) &= \frac{\mathrm{d}}{\mathrm{d}t}\left(e^{\frac{1}{2}t^2}\right) = t e^{\frac{1}{2}t^2}, \\
M_X^{(2)}(t) &= \frac{\mathrm{d}}{\mathrm{d}t}\left(t e^{\frac{1}{2}t^2}\right) = (t^2 + 1) e^{\frac{1}{2}t^2}, \\
M_X^{(3)}(t) &= t(t^2 + 3) e^{\frac{1}{2}t^2}, \\
M_X^{(4)}(t) &= (t^4 + 6t^2 + 3) e^{\frac{1}{2}t^2}, \\
&\vdots
\end{aligned}$$

Theorem 71.
$$M_{aX+b}(t) = e^{bt} M_X(at).$$

Proof.
$$M_{aX+b}(t) = E\left[e^{t(aX+b)}\right] = E\left[e^{atX} e^{bt}\right] = E\left[e^{atX}\right] e^{bt} = e^{bt} M_X(at).$$
□

Example 72. Suppose $Z \sim \mathrm{N}(0,1)$ and $X = \mu + Z$. Then $X \sim \mathrm{N}(\mu, \sigma^2)$.
$$M_X(t) = e^{\mu t} M_Z(\sigma t) = e^{\mu t} e^{\frac{\sigma^2 t^2}{2}} = e^{\mu t + \frac{\sigma^2 t^2}{2}}.$$

Theorem 73. Let $X_1, \ldots, X_n$ be a sequence of *independent* random variables with MGFs $M_{X_1}(t), \ldots, M_{X_n}(t)$, and let $Z = X_1 + \cdots + X_n$. Then
$$M_Z(t) = \prod_{i=1}^n M_{X_i}(t).$$

Proof.
$$M_Z(t) = E\left[e^{t(X_1+\cdots+X_n)}\right] = E\left[\prod_{i=1}^n e^{X_i t}\right] = \prod_{i=1}^n E\left[e^{X_i t}\right] = \prod_{i=1}^n M_{X_i}(t).$$
□

Example 74. Suppose $X_i \sim \mathrm{Gamma}(\alpha_i, \beta)$ for $i = 1, \ldots, n$ are i.i.d., and $S = \sum_{i=1}^n X_i$. Then
$$M_{X_i}(t) = \left(\frac{\beta}{\beta - t}\right)^{\alpha_i} \quad \text{so that} \quad M_S(t) = \left(\frac{\beta}{\beta - t}\right)^{\sum_{i=1}^n \alpha_i}$$
and so $S \sim \mathrm{Gamma}\left(\sum_{i=1}^n \alpha_i, \beta\right)$.

Theorem 75. Let $X_1, \ldots, X_n$ be i.i.d. random variables, each with MGF, $M_X(t)$. Then
$$M_{\overline{X}}(t) = \left[M_X\left(\frac{t}{n}\right)\right]^n.$$

Proof. Write $\overline{X} = \frac{1}{n}X_1 + \frac{1}{n}X_2 + \cdots + \frac{1}{n}X_n$, and then apply Theorems 71 and 73. □

Theorem 76. If $M_X(t)$ exists (in a neighborhood of zero), then for $r = 0, 1, 2, \ldots$

(i) $M_X^{(r)}(t)$ exists near zero, and

(ii) $E\left(|X^r|\right) < \infty$.

(Characterization) Theorem 77. If the MGFs of X and Y exist and $M_X(t) = M_Y(t)$ in a neighborhood of zero, then

$$F_X(u) = F_Y(u) \quad \forall u,$$

i.e. MGFs characterize distributions of random variables.

Example 78. Refer to the Gamma distribution example, which is Example 74.

Theorem 79. Let $X_1, X_2, \ldots$ be a countable sequence of random variables with MGFs $M_{X_1}(t), M_{X_2}(t), \ldots$ so that

$$\lim_{i\to\infty} M_{X_i}(t) = M_X(t) \quad \text{for } t \text{ in a neighborhood of zero,}$$

where $M_X(t)$ is an MGF. Then there is a unique cdf, F_X, for which

$$\lim_{i\to\infty} F_{X_i}(x) = F_X(x) \quad \text{for all } x \text{ where } F_X(x) \text{ is continuous,}$$

that is, $X_n \to_{\mathrm{d}} X$. The moments of $F_X(x)$ are determined by $M_X(t)$.

Comments

- Proofs of Theorems 76, 77, and 79 follow from the properties of Laplace transforms, which are beyond the scope of the course.
- We can now show convergence in distribution by showing convergence of MGF in a neighborhood of zero.
- Convergence of MGF is a sufficient, but *not* necessary condition for convergence in distribution. (The MGF may not exist in a neighborhood of zero.)
- A more general strategy involves characteristic functions, $\phi_X(t) = E\left(e^{itX}\right)$.
- Moments alone do not characterize distributions. That is, there exists X and Y s.t. $E(X^r) = E(Y^r)$ for $r = 0, 1, 2, \ldots$, but $F_X \neq F_Y$.
  - However, if X and Y have finite support, and all moments exist, then $F_X(u) = F_Y(u)$ for all u iff $E(X^r) = E(Y^r)$ for $r = 0, 1, 2, \ldots$.

7.3 Central Limit Theorem

(Central Limit Theorem) Theorem 80. Let $X_1, X_2, \ldots, X_n$ be a sequence of i.i.d. random variables with $E(X_i) = \mu$ and $\mathrm{Var}(X_i) = \sigma^2 < \infty$, then

$$\lim_{n\to\infty} \frac{\overline{X} - \mu}{\frac{\sigma}{\sqrt{n}}} \sim \mathrm{N}(0,1) \quad \text{or} \quad \frac{\overline{X} - \mu}{\frac{\sigma}{\sqrt{n}}} \to_{\mathrm{d}} \mathrm{N}(0,1).$$

Equivalently, when n is large we have approximately

$$\overline{X} \sim \mathrm{N}\left(\mu, \frac{\sigma^2}{n}\right) \quad \text{or} \quad \sum_{i=1}^{n} X_i \sim \mathrm{N}(n\mu, n\sigma^2).$$

Proof. By assumption, $M_X(t)$ exists in a neighborhood of zero s.t. $E(X_i) = \mu$ and $\mathrm{Var}(X_i) = \sigma^2$ exists. Let $Z_n = \dfrac{\overline{X}_n - \mu}{\frac{\sigma}{\sqrt{n}}}$. It is sufficient to show that

$$\lim_{n\to\infty} M_{Z_n}(t) = M_Z(t) = e^{\frac{t^2}{2}},$$

where $Z \sim \mathrm{N}(0,1)$. Re-write Z_n as

$$Z_n = \frac{X_1 + X_2 + \cdots + X_n - n\mu}{\sigma\sqrt{n}} = \frac{1}{\sqrt{n}}\left(\frac{X_1 - \mu}{\sigma} + \frac{X_2 - \mu}{\sigma} + \cdots + \frac{X_n - \mu}{\sigma}\right).$$

Define $W_i = \dfrac{X_i - \mu}{\sigma}$ s.t. $E(W_i) = 0, \mathrm{Var}(W_i) = 1$, and we can rewrite Z_n as

$$Z_n = \frac{1}{\sqrt{n}}\left(W_1 + W_2 + \cdots + W_n\right).$$

Therefore,

$$\begin{aligned} M_{W_i}(t) = E\left(e^{tW_i}\right) &= M_{W_i}(0) + M_{W_i}^{(1)}(0) + M_{W_i}^{(1)}(0)t + \frac{M_{W_i}^{(2)}(0)}{2!}t^2 + O(t^3) \\ &= 1 + (0)t + \frac{t^2}{2} + O(t^3) \\ &= 1 + \frac{t^2}{2} + O(t^3). \end{aligned}$$

Thus,

$$M_{\frac{W_i}{\sqrt{n}}}(t) = M_{W_i}\left(\frac{t}{\sqrt{n}}\right) = 1 + \frac{t^2}{2n} + O\left(\left(\frac{t}{\sqrt{n}}\right)^3\right).$$

$$\Longrightarrow M_{Z_n}(t) = \prod_{i=1}^{n} M_{\frac{W_i}{\sqrt{n}}}(t) = \left(1 + \frac{t^2}{2n} + O\left(\left(\frac{t}{\sqrt{n}}\right)^3\right)\right)^n.$$

By taking $\log[M_{Z_n}(t)]$, and the taylor expansion

$$\log(1+x) = x - \frac{x^2}{2} + \frac{x^3}{3} - \ldots = x + O(x^2),$$

we have

$$\log[M_{Z_n}(t)] = n\left[\frac{t^2}{2n} + O\left(\left(\frac{t}{\sqrt{n}}\right)^3\right) + O\left(\frac{t^4}{n^2}\right)\right] = \frac{t^2}{2} + O\left(\frac{t^3}{n^{\frac{1}{2}}}\right) + O\left(\frac{t^4}{n}\right) \longrightarrow \frac{t^2}{2}$$

as $n \to \infty$. Thus,

$$\lim_{n\to\infty} M_{Z_n}(t) = e^{\frac{t^2}{2}} = M_Z(t), \quad Z \sim \mathrm{N}(0,1),$$

i.e.

$$Z_n = \frac{\overline{X}_n - \mu}{\frac{\sigma}{\sqrt{n}}} \to_{\mathrm{d}} \mathrm{N}(0,1).$$

□

7.4 The Law of Large Numbers and Inequalities

Theorem 81. Let X be a random variable and $g(\cdot)$ be a non-negative function. Then $\forall r > 0$,

$$P(g(X) \geq r) \leq \frac{E[g(X)]}{r}.$$

Proof.

$$\begin{aligned} E(g(X)) &= \int_{-\infty}^{\infty} g(x) f_X(x)\mathrm{d}x \\ &\geq \int_{\{x:g(x)\geq r\}} g(x) f_X(x)\mathrm{d}x \\ &\geq \int_{\{x:g(x)\geq r\}} r f_X(x)\mathrm{d}x \\ &= r\int_{\{x:g(x)\geq r\}} f_X(x)\mathrm{d}x = rP(g(X) \geq r). \end{aligned}$$

□

Theorem 82. Suppose $X_1, X_2, \ldots$ is a sequence of i.i.d. random variables with $E(X_i) = \mu$ and $\mathrm{Var}(X_i) = \sigma^2 < \infty$. If $\overline{X}_n = \frac{1}{n}\sum_{i=1}^{n} X_i$, then $\overline{X}_n \to_{\mathrm{p}} \mu$, i.e.

$$\forall \epsilon > 0, \lim_{n\to\infty} P(|\overline{X}_n - \mu| < \epsilon) = 1.$$

Proof. Apply Chebyshev's inequality to random variable $\overline{X}_n$. Let $g(x) = (x-\mu)^2$ s.t. $E[g(\overline{X})] = E[(\overline{X}-\mu)^2] = \mathrm{Var}(\overline{X}) = \frac{\sigma^2}{n}$. Therefore,

$$P(|\overline{X}_n - \mu| \geq \epsilon) = P((\overline{X}_n - \mu)^2 \geq \epsilon^2) \leq \frac{E[(\overline{X}_n - \mu)^2]}{\epsilon^2} = \frac{\mathrm{Var}(\overline{X}_n)}{\epsilon^2} = \frac{\sigma^2}{n\epsilon^2} \longrightarrow 0$$

as $n \to \infty$, therefore $P(|\overline{X}_n - \mu| < \epsilon) \longrightarrow 1$. □

(Jensen's Inequality) Theorem 83.

- If $g(x)$ is a *convex* function ($g'' \geq 0$), then $E[g(X)] \geq g[E(X)]$.
- If $g(x)$ is a *concave* function ($g'' \leq 0$), then $E[g(X)] \leq g[E(X)]$.

Example 84. Suppose $X \sim \text{Binomial}(n, p)$ and consider the estimator $\hat{P} = \frac{X}{n}$ of p. We can show that $\hat{P}$ is an unbiased estimator of p by

$$E(\hat{P}) = E\left(\frac{X}{n}\right) = \frac{E(X)}{n} = \frac{np}{n} = p.$$

Suppose we want an estimator of the odds ratio $\xi = \frac{p}{1-p} = g(p)$.

$$\begin{aligned} g'(p) &= \frac{1-p+p}{(1-p)^2} = \frac{1}{(1-p)^2}, \\ g''(p) &= \frac{2}{(1-p)^3} > 0 \quad \text{if } p \in (0,1). \end{aligned}$$

Consider the estimator $\Xi = \frac{\hat{P}}{1-\hat{P}}$ of ξ,

$$E(\Xi) = E[g(\hat{P})] \geq g[E(\hat{P})] = g(p) = \xi,$$

implying that Ξ is biased.

In general if we have an unbiased estimator $\hat{\theta}$ of parameter θ and want to estimate some function of the parameter $\phi = g(\theta)$ using the estimator $\hat{\phi} = g(\hat{\theta})$ it is important to realize that

$$E(\hat{\phi}) = E[g(\hat{\theta})] \neq g[E(\hat{\theta})] = g(\theta) = \phi,$$

i.e. unbiasedness is *not* necessarily invariant to transformation.

Chapter 8

Estimation

8.1 Estimators

Definition 85.

- A ***statistic*** is a function $T = T(X_1, X_2, \ldots, X_n) = T(\mathbf{X})$, and *is* itself a random variable.
- If a statistic $T(\mathbf{X})$ is to be used to approximate parameters of the distribution $P_{\mathbf{X}|\theta}(\cdot)$, we say that T is an ***estimator*** for those parameters, and we call the actual realized value of the estimator for a particular data sample, $t(\mathbf{x})$, an ***estimate***.
- A ***point estimate*** is a statistic estimating a single parameter or characteristic of a distribution.

Definition 86. We define the ***bias*** of an estimator T for a parameter $g(\theta)$ as

$$\text{bias}_\theta(T) = E_\theta(T) - g(\theta).$$

If the estimator has zero bias, we say the estimator is ***unbiased***.

Definition 87. We define the ***bias-corrected sample variance*** as

$$S_{n-1}^2 = \frac{1}{n-1} \sum_{i=1}^{\infty} \left(X_i - \overline{X}\right)^2,$$

which is then always an unbiased estimator of the population variance σ^2.

Definition 88. Suppose we have two unbiased estimators for a parameter θ, which we call $\hat{\Theta}(\mathbf{X})$ and $\hat{\Psi}(\mathbf{X})$. We say $\hat{\Theta}$ is ***more efficient*** than $\hat{\Psi}$ if

1. $\forall \theta, \mathrm{Var}_\theta(\hat{\Theta}) \leq \mathrm{Var}_\theta(\hat{\Psi})$,
2. $\exists \theta$ s.t. $\mathrm{Var}_\theta(\hat{\Theta}) < \mathrm{Var}_\theta(\hat{\Psi})$.

If $\hat{\Theta}$ is more efficient than any other possible estimator, we say $\hat{\Theta}$ is ***efficient***.

Definition 89. We say that $\hat{\Theta}$ is a ***consistent*** estimator for the parameter θ if $\hat{\Theta}$ *converges in probability* to θ, i.e.

$$\forall \epsilon > 0, P(|\hat{\Theta} - \theta| > \epsilon) \to 0 \text{ as } n \to \infty.$$

This is hard to demonstrate, but if $\hat{\Theta}$ is unbiased we do have

$$\lim_{n\to\infty} \mathrm{Var}(\hat{\Theta}) = 0 \Longrightarrow \hat{\Theta} \text{ is consistent.}$$

Example 90. Let $\overline{X}$ be the estimator for the population mean μ. We can show that $\overline{X}$ is unbiased, efficient, and consistent!

8.2 Maximum Likelihood Estimation

Definition 91. Let $\theta \in \Theta$ be a parameter of a population where Θ is the parameter space, and $\mathbf{x} \in \mathbb{R}^n$ be the realization of the random object $\mathbf{X} \in \mathbb{R}^n$, $n \in \mathbb{Z}^+$. The ***likelihood function*** of $\mathbf{x}$ is

$$L(\theta) = L(\theta|\mathbf{x}) = \begin{cases} P_{\mathbf{X}|\theta}(\mathbf{X} = \mathbf{x}), & \text{discrete data,} \\ f_{\mathbf{X}|\theta}(\mathbf{x}), & \text{absolutely continuous data.} \end{cases}$$

Then the ***maximum likelihood estimator (MLE)*** of θ is an estimator $\hat{\theta}$ s.t.

$$L(\hat{\theta}) = \sup_{\theta \in \Theta} L(\theta).$$

$\hat{\theta}$ could be obtained by

1. find out $L(\theta)$.
2. take $\log(L(\theta))$ to obtain the *log-likelihood function* $l(\theta)$.
3. partial differentiate $l(\theta)$ w.r.t. θ and equate $l(\theta)$ to 0.
4. solve for θ to obtain the expression for $\hat{\theta}$.
5. check that the obtained $\hat{\theta}$ corresponds to a maximum of the likelihood function by checking if
$$\frac{\partial^2}{\partial \theta^2} l(\hat{\theta}) < 0.$$
6. Change from $\hat{\theta} = \hat{\theta}(x)$ to $\hat{\theta} = \hat{\theta}(X)$ so that the final expressions is a *estimator* instead an *estimate.*

Comments

− The MLE is not necessarily unbiased, i.e. it is possible that
$$\hat{\theta} \neq \theta,$$
where $\hat{\theta}$ is the MLE and θ is the true parameter.

\+ The MLE is consistent.

\+ The MLE is asymptotically normal, i.e. assume $(\hat{\theta}_n)_{n \in \mathbb{N}}$ is a consistent sequence of MLEs, then
$$\sqrt{n}(\hat{\theta}_n - \theta) \to_{\mathrm{d}} N(0, \sigma^2(\theta))$$
for some $\sigma^2(\theta)$.

\+ The MLE is always asymptotically efficient, and if an efficient estimator exists, it is the MLE.

8.3 Confidence Intervals

Definition 92. In general, a $1-\alpha$ ***confidence interval*** I is a random interval that contains the "true" parameter with probability $\geq 1-\alpha$, i.e.

$$P_\theta(\theta \in I) \geq 1-\alpha, \quad \forall \theta \in \Theta.$$

If sample size is large, we can exploit the CLT. Thus, for any desired coverage probability level $1-\alpha$ we can define the $1-\alpha$ C.I. for θ by

$$\left[\hat{\theta} - z_{1-\frac{\alpha}{2}}\frac{\sigma}{\sqrt{n}}, \hat{\theta} + z_{1-\frac{\alpha}{2}}\frac{\sigma}{\sqrt{n}}\right],$$

where z_α is the α-quantile of the standard normal.

Example 93. A corporation conducts a survey to investigate the proportion of employees who thought the board was doing a good job. 1000 employees, randomly selected, were asked, and 732 said they did. Find a 99% confidence interval for the value of the proportion in the population who thought the board was doing a good job.

We can model each observation as $X_i \sim \text{Bernoulli}(p)$ for some unknown p, and we want to find a C.I. for p, which is also the mean of X. We have our estimate $\hat{p} = \overline{x} = 0.732$ for which we use the CLT. Since the variance of Bernoulli(p) is $p(1-p)$, we can use $\overline{x}(1-\overline{x}) = 0.196$ as an approximate variance. So an approximate 99% C.I. is

$$\left[0.732 - 2.576 \times \sqrt{\frac{0.196}{1000}},\ 0.732 + 2.576 \times \sqrt{\frac{0.196}{1000}}\right].$$

Theorem 94. If $X_1, X_2, \ldots, X_n$ are an i.i.d. sample from $\mathrm{N}(\mu, \sigma^2)$ with σ known, then we can construct an *exact* confidence interval for μ as

$$\left[\overline{x} - z_{1-\frac{}{2}}\frac{\sigma}{\sqrt{n}}, \overline{x} + z_{1-\frac{\alpha}{2}}\frac{\sigma}{\sqrt{n}}\right].$$

If σ is unknown, then we can construct an C.I. for μ as

$$\left[\overline{x} - t_{n-1,1-\frac{\alpha}{2}}\frac{s_{n-1}}{\sqrt{n}},\ \overline{x} + t_{n-1,1-\frac{\alpha}{2}}\frac{s_{n-1}}{\sqrt{n}}\right],$$

where $t_{\nu,\alpha}$ is the α-quantile of Student(ν), which is the Student's t-distribution with ν degrees of freedom, and $S_{n-1} = \sqrt{S_{n-1}^2}$ is the bias-corrected sample standard deviation.

Comments

- Student(ν) is heavier tailed than $\mathrm{N}(0,1)$ for any number of degrees of freedom ν, so the t-distribution C.I. will always be wider than the Normal distribution C.I. Therefore if we know σ^2, we should use it.

- $\lim_{\nu \to \infty} \text{Student}(\nu) = \text{N}(0, 1)$.
- For $\nu > 40$, the difference between Student(ν) and N$(0, 1)$ is so insignificant that the t-distribution is not tabulated beyond this many degrees of freedom, and so there we can instead revert to N$(0, 1)$ tables.

Chapter 9

Hypothesis Testing

9.1 Definitions

Definition 95. We partition the parameter space Θ into two disjoint sets Θ_0 and Θ_1 and we wish to test

$$H_0 : \theta \in \Theta_0 \quad \text{v.s.} \quad H_1 : \theta \in \Theta_1.$$

We call the H_0 the ***null hypothesis*** and H_1 the ***alternative hypothesis***.

To test the validity of H_0, we first choose a test statistic $T(\mathbf{X})$ of the data for which we can find the distribution P_T under H_0. Then we identify a rejection region $R \subset \mathbb{R}$ of low probability values of T under the assumption that H_0 is true, i.e. a region R s.t.

$$P(T \in R | H_0) = \alpha$$

for some ***significance level*** $\alpha \in (0, 1)$. We finally calculate the observed test statistic $t(\mathbf{x})$ and

- if $t \in R$ we "reject the null hypothesis at the α level".
- if $t \in R$ we "do not reject (retain) the null hypothesis at the α level".

Definition 96. We define the ***p-value*** as

$$p = \sup_{\theta \in \Theta_0} P_\theta(\text{observing sth "at least as extreme" as the observation}).$$

Definition 97. The ***power function*** is defined as the mapping

$$\beta : \Theta \mapsto [0,1].$$

The ***power*** of a hypothesis test is then defined as

$$\beta(\theta) = P_\theta(\text{reject } H_0).$$

If $\theta \in \Theta_0$ then we want $\beta(\theta)$ to be small; If $\theta \in \Theta_1$ then we want $\beta(\theta)$ to be large.

The following Figure 9.1 defines the ***Type I error*** and ***Type II error***.

	H_0 true	H_0 false
do not reject H_0	✓	Type II error
reject H_0	Type I error	✓

False negative

False positive

power function

[0,1]

Figure 9.1: Type I and Type II errors

Example 98. $X \sim \mathrm{N}(\theta, 1), \theta \in \mathbb{R}$ unknown. We test

$$H_0 : \theta \leq 0 \quad \text{against} \quad H_1 : \theta > 0.$$

Thus $\Theta = \mathbb{R}$, $\Theta_0 = (-\infty, 0]$, $\Theta_1 = (0, \infty)$. Suppose we use the critical (rejection) region

$$R = [c, \infty).$$

We choose a critical value c s.t. the test is of level α. For $\theta \leq 0$:

$$P_\theta(\text{reject } H_0) = P_\theta(X \geq c) = P_\theta(\underbrace{X - \theta}_{\sim \mathrm{N}(0,1)} > c - \theta) = 1 - \Phi(c - \theta) \leq 1 - \Phi(c).$$

Thus we choose c s.t. $\Phi(c) = 1 - \alpha$, then $\forall\, \theta \in \Theta_0, P_\theta(\text{reject } H_0) \leq \alpha$.

9.2 Testing for a Population Mean

Suppose $X_1, X_2, \ldots, X_n$ are i.i.d. $\mathrm{N}(\mu, \sigma^2)$. We wish to test if $\mu = \mu_0$ for some specific value μ_0, so we can state our null and alternative hypothesis as

$$H_0 : \mu = \mu_0 \quad \text{v.s.} \quad H_1 : \mu \neq \mu_0$$

9.2.1 Normal Distribution with Known Variance

Say σ^2 is known and μ is unknown. We can set up a test statistic

$$Z = \frac{\overline{X} - \mu_0}{\frac{\sigma}{\sqrt{n}}} \sim \Phi.$$

Thus the rejection region R can be defined as

$$R = \left(-\infty, -z_{1-\frac{\alpha}{2}}\right) \cup \left(z_{1-\frac{\alpha}{2}}, \infty\right) = \left\{z \middle| |z| > z_{1-\frac{\alpha}{2}}\right\}$$

s.t. $P(Z \in R|H_0) = \alpha$. As such, we reject H_0 at the α significance level $\iff$ our observed test statistic satisfies

$$z = \frac{\overline{x} - \mu_0}{\frac{\sigma}{\sqrt{n}}} \in R \quad \text{or} \quad p\text{-value} = 2\left(1 - \Phi(|z|)\right) \leq \alpha.$$

9.2.2 Normal Distribution with Unknown Variance

Say σ^2 is unknown and μ is unknown. We can set up a test statistic

$$T = \frac{\overline{X} - \mu_0}{\frac{s_{n-1}}{\sqrt{n}}} \sim t_{n-1}.$$

The rejection region R thus changes to

$$R = \left\{t \middle| |t| > t_{n-1,1-\frac{\alpha}{2}}\right\}$$

s.t. $P(T \in R|H_0) = \alpha$. As such, we reject H_0 at the α significance level $\iff$ our observed test statistic satisfies

$$t = \frac{\overline{x} - \mu_0}{\frac{s_{n-1}}{\sqrt{n}}} \in R \quad \text{or} \quad p\text{-value} = 2\left(1 - t_{n-1,1-\frac{\alpha}{2}}\right) \leq \alpha.$$

9.3 Testing for Differences in Population Means

Suppose that

- $\mathbf{X} = (X_1, \ldots, X_{n_1})$ are i.i.d. $\mathrm{N}(\mu_X, \sigma_X^2)$ with μ_X unknown;
- $\mathbf{Y} = (Y_1, \ldots, Y_{n_1})$ are i.i.d. $\mathrm{N}(\mu_Y, \sigma_Y^2)$ with μ_Y unknown;
- the two samples $\mathbf{X}$ and $\mathbf{Y}$ are independent,

and we want to test

$$H_0 : \mu_X = \mu_Y \quad \text{v.s.} \quad H_1 : \mu_X \neq \mu_Y.$$

9.3.1 Normal Distribution with Known Variances

If σ is known, we have

$$\overline{X} \sim \mathrm{N}\left(\mu_X, \frac{\sigma_X^2}{n_1}\right), \quad \overline{Y} \sim \mathrm{N}\left(\mu_Y, \frac{\sigma_Y^2}{n_1}\right) \quad \Longrightarrow \quad \overline{X} - \overline{Y} \sim \mathrm{N}\left(\mu_X - \mu_Y, \frac{\sigma_X^2}{n_1} + \frac{\sigma_Y^2}{n_2}\right),$$

thereby setting up test statistic

$$Z = \frac{(\overline{X} - \overline{Y}) - (\mu_X - \mu_Y)}{\sqrt{\frac{\sigma_X^2}{n_1} + \frac{\sigma_Y^2}{n_2}}} = \frac{(\overline{X} - \overline{Y})}{\sqrt{\frac{\sigma_X^2}{n_1} + \frac{\sigma_Y^2}{n_2}}} \sim \Phi,$$

following up with the investigation of rejection of null hypothesis.

9.3.2 Normal Distribution with Unknown Variances

If σ is unknown, we have

$$T = \frac{(\overline{X} - \overline{Y}) - (\mu_X - \mu_Y)}{S_{n_1+n_2-2}\sqrt{\frac{1}{n_1} + \frac{1}{n_2}}} = \frac{(\overline{X} - \overline{Y})}{S_{n_1+n_2-2}\sqrt{\frac{1}{n_1} + \frac{1}{n_2}}} \sim t_{n_1+n_2-2},$$

following up with the investigation of rejection of null hypothesis.

9.4 Goodness of Fit

9.4.1 Chi-square Test

Definition 99. To test for ***goodness of fit***, i.e. compare the *observed frequency* $\mathbf{O} = (O_1, \dots, O_k)$ with the *expected frequency* $\mathbb{E} = (E_1, \dots, E_k)$, we set up $H_0 : \theta = \theta_0$ v.s. $H_1 : \theta \neq \theta_0$ for the value of the unknown parameter, and use the ***chi-square statistic***

$$\chi^2 = \sum_{i=1}^{k} \frac{(O_i - E_i)^2}{E_i}. \quad (\geq 0)$$

If H_0 were true, thenthe statistic χ^2 would approximately follow a ***chi-square distribution*** with $\nu = k - m - 1$ degrees of freedom.

Comments

- k is the number of values (categories) the simple random variable X can take.
- m is the number of parameters we needed to estimate from the data ($\dim(\theta)$) in order to calculate the p_j's.

  - E.g. given a sample without specifying the model, the degree of freedom $\nu = k - 0 - 1 = k - 1$; if say the sample is fitted with a Poisson distribution with rate parameter λ, then $\nu = k - 1 - 1 = k - 2$.

- For the approximation to be valid, we should have $\forall j, E_j \geq 5$. This may require some merging of categories.

- Larger χ^2 corresponds to larger deviations from the null hypothesis model; if $\chi^2 = 0$, observed counts exactly match those expected under H_0.

- Since $\chi^2 \geq 0$, we always perform a one-sided goodness of fit test using the χ^2 statistic, looking at the upper tail of the distribution, leading to the rejection region R at α level being
$$R = \left\{ x^2 | x^2 > \chi^2_{k-m-1,1-\alpha} \right\}.$$

9.4.2 Independence using Chi-square Statistic

Assume two discrete random variables X and Y that can each take finite values which are jointly distributed with unknown probability mass function p_{XY}. To determine if X and Y are independent, we can do the following:

Let the ranges of X and Y be $\{x_1, \dots, x_k\}$ and $\{y_1, \dots, y_l\}$ respectively. Then we can form the following $k \times l$ ***contingency table***:

Table 9.1: contingency table example

$$\begin{array}{c|cccc|c}
 & y_1 & y_2 & \cdots & y_l & \\
\hline
x_1 & n_{11} & n_{12} & & n_{1l} & n_{1\bullet} \\
x_2 & n_{21} & n_{22} & & n_{2l} & n_{2\bullet} \\
\vdots & & & & & \\
x_k & n_{k1} & n_{k2} & & n_{kl} & n_{k\bullet} \\
\hline
 & n_{\bullet 1} & n_{\bullet 2} & \cdots & n_{\bullet l} & n
\end{array}$$

where n_{ij} represents the number of times we observe the pair (x_i, y_i), $n_{i\bullet}$ represents the frequencies of x_i in the sample, and similarly for $n_{\bullet j}$. Under the null hypothesis,

$$H_0 : X \text{ and } Y \text{ are independent},$$

we can compute the expected values of entries of the contingency table by

$$\hat{n}_{ij} = \frac{n_{i\bullet} n_{\bullet j}}{n}$$

since

$$\hat{p}_{i\bullet} = p_X(x_i) = \frac{n_{i\bullet}}{n}, \quad \hat{p}_{\bullet j} = p_Y(y_j) = \frac{n_{\bullet j}}{n} \quad \Longrightarrow \quad \hat{p}_{ij} = \hat{p}_{i\bullet} \times \hat{p}_{\bullet j} = \frac{n_{i\bullet} n_{\bullet j}}{n^2}$$

and multiply boths sides with n to obtain the desired quantity $\hat{n}_{ij}$. We can then set up the chi-square test statistic

$$x^2 = \sum_{i,j} \frac{(n_{ij} - \hat{n}_{ij})^2}{\hat{n}_{ij}}$$

with the degress of freedom $\nu = kl - (k-1) - (l-1) - 1 = (k-1)(l-1)$. Hence the rejection region for a hypothesis test of independence in a $k \times l$ contingency table at α level is given by

$$R = \left\{ x^2 | x^2 > \chi^2_{(k-1)(l-1),1-\alpha} \right\}.$$